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Do Budget Deficits Matter? Some Pre-World War II Evidence

W. Douglas McMillin and Thomas R. Beard

We estimate five-variable vector autoregressions comprising deficits, money, prices, industrial production, and interest rates. The impact of budget deficits is evaluated by variance decompositions. In estimating vector autoregressions for the July 1922–June 1938 period we first use a technique in which lag lengths differ for each variable in each equation and for the same variable across equations. We next estimate vector autoregressions using a common lag length for the same time period and for two subperiods. While the impact of the deficit varies, there is no evidence of substantial debt monetization or of important effects on the other variables.

I. Introduction

A substantial literature on the impact of federal budget deficits on the U.S. money supply, interest rates, and other macro variables focuses heavily on the post–World War II period. Unfortunately, the empirical results have been mixed. Numerous studies found evidence that the Federal Reserve monetizes budget deficits—e.g., Hamburger and Zwick (1981) and Allen and Smith (1983)—while other studies suggested no monetization—e.g., Barro (1978) and Niskanen (1978). Likewise, evidence on the impact of deficits on interest rates and on real and nominal output is mixed. For example, Makin (1983) reported a positive effect of deficits on short-term rates and Hoelscher (1986) found strong positive effects on long-term rates. Another study by Hoelscher (1983) found no significant effect on short-term rates while Mascaro and Meltzer (1983) found no significant effects on either long-term or short-term rates. While a majority of studies have found no significant connection between deficits and interest rates, the empirical results appear to be sensitive to the choice of dependent and independent variables, the way the deficit is measured, and the time period examined (Barth et al. 1984).

A number of studies have estimated vector autoregressions (VARs) and used Granger-

The authors are Associate Professor and Professor, respectively, Louisiana State University, Baton Rouge, Louisiana.

Address reprint requests to Thomas R. Beard, Department of Economics, Louisiana State University, 2107 CEBA, Baton Rouge, Louisiana 70803-6306.

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causality tests in studying the impact of budget deficits. For example, Dwyer (1982), using quarterly data for 1952–1978, found that government debt acquired by the private sector did not Granger-cause debt acquired by the Federal Reserve or the money supply (or, for that matter, any other macro variable). McMillin (1986a), however, found that several measures of the deficit Granger-caused the monetary base for the 1961:1–1979:3 period. Mixed evidence was found by Thornton (1984) for several post-World War II periods. In an important and carefully constructed paper, Plosser (1982) found that a substitution of debt for tax financing of a given level of government expenditures did not increase interest rates, although an increase in government purchases did raise rates. His data covered the period 1954:1–1978:4. Using quarterly data for 1957–1984, McMillin (1986b) found four deficit measures did not Granger-cause short-term rates, while Canto and Rapp (1982), using annual data for 1929–1980, found no evidence that two deficit measures caused interest rates. Blinder (1983), using quarterly data for fiscal years 1952–1981, found that deficits help to predict inflation but not real output. Miller (1983), on the other hand, found substantial effects of deficits (understood as different deficit "policies") on output, inflation, and interest rates over four post-World War II subperiods beginning in 1948. For recent surveys or reviews of the deficit literature, see Dwyer (1985), Beard and McMillin (1986), U.S. Treasury (1984), Congressional Budget Office (1984), and Seater (1985).

That empirical results have not been more consistent is perhaps not surprising in view of the theoretical ambiguity about the macroeconomic effects of budget deficits. The conventional view that federal debt is a component of private sector wealth and that deficits thus increase wealth and alter macro variables stands in contrast to the Ricardian equivalence hypothesis as developed by Barro (1974). According to this view, an increase in government debt is equivalent to a future increase in taxes and thus is not an addition to private wealth. A switch from lump sum tax finance of a given level of government purchases to debt finance has no effect on consumption, interest rates, or aggregate demand. However, even in the Ricardian framework, changes in government purchases and in distortionary, or nonlump sum, taxes have real effects.

The Ricardian equivalence hypothesis has not been universally accepted. Buiter and Tobin (1979), for example, have argued against its underlying assumptions and concluded that tax and debt finance are not equivalent. In a detailed survey of more recent literature, Brunner (1986) explored the conditions under which equivalence may not hold. Thus, there is no theoretical consensus on the macroeconomic role of government deficits and debt.

Despite the importance of the budget deficit question, little attention has been devoted to the period prior to World War II, and especially to the period between the two world wars. In addition to being relatively neglected, the interwar period is also of interest in that the government ran surpluses as well as deficits in various years, in contrast to the persistent deficits since World War II. The most common explanation for debt monetization assigns the Federal Reserve a crucial institutional link in the monetization process. Since the reaction functions of the pre- and post-World War II Federal Reserve are possibly quite different, it is useful to investigate whether the government's lending and borrowing activities in this earlier period affected interest rates (and other variables) and alleviated or created pressures for debt monetization.

Both Joines (1985) and Evans (1987) have studied the impact of budget deficits over a long time span that includes the interwar years; Joines found no evidence of debt monetization and Evans found no evidence that deficits affect interest rates.[1] In this paper, we focus on the interwar period. Both Sims (1980) and Burbidge and Harrison (1985) have studied the interwar period, although neither considered a fiscal policy variable in their models. Both used monthly data in estimating four-variable VARs comprising money, prices, industrial production, and interest rate variables. We focus on budget deficits and estimate five-variable VARs comprising deficits, money, prices, industrial production, and interest rates. The impact of budget deficits is evaluated by Granger-causality tests and by variance decompositions (VDCs) based on the estimated VARs.

An important consideration in studying the interwar years is the choice of time period. In both Sims and Burbidge and Harrison the interwar period refers to January 1920–December 1941, with data from 1919 used for initial conditions. In our view, it is desirable to select a shorter time period that focuses exclusively on a peacetime economy. Firestone (1960), analyzing monthly data, identified the war cycles (measured trough to trough) as December 1914–March 1919 and June 1938–October 1945. He also identified a postwar cycle as March 1919–July 1921. The first of his four interwar peacetime cycles thus began in July 1921 and the fourth ended in June 1938.

In this paper, we define the interwar period as corresponding to the four peacetime cycles in Firestone. This avoids any problems connected with war-related government expenditures. With a maximum lag length of 12 months considered, no data prior to July 1921 are used for initial conditions.

In section II our basic methodology is discussed; the VAR technique we use results in a system in which all variables enter each equation in the system, but, in general, the lag lengths will differ for each variable in each equation and for the same variable across equations. Empirical results for this technique are reported in section III. In section IV we examine the results of an alternative specification of the VAR that employs the same lag length for each variable in each equation, as was done by Sims and Burbidge and Harrison. Our conclusions are reported in section V.

II. Methodology

The methodology employed to analyze the macroeconomic effects of federal deficits is a variant of the vector autoregressive technique suggested by Hsiao (1979, 1981) and extended by Caines, Keng, and Sethi (1981) and Keng (1982). This technique rather than a structural model is used in order to avoid imposing potentially spurious a priori constraints on the model (such as, for example, econometric exogeneity of the deficit in the industrial production equation). Rather than employing the same lag length for each variable in each equation, as was done by Sims (1980) or Burbidge and Harrison (1985),

[1] Using a traditional reaction function technique and yearly data, Joines (1985) examined the relationship between deficits and the growth of the monetary base over a very long period, 1872–1983, and over several subperiods, including 1915–1953. An initial equation showed a positive relation between deficits and monetary base growth over this subperiod; this relationship disappeared, however, when lagged unemployment was added to the model. Evans (1987), using regression techniques and monthly data, examined the relationship between deficits and interest rates over the period June 1908–March 1984, and over eleven subperiods including January 1920–December 1929 and January 1930–December 1939. He found no statistically significant effect between deficits and interest rates in these subperiods. Neither Joines nor Evans looked specifically at the interwar period as a whole.

Akaike's final prediction error (*FPE*) criterion is used to determine the appropriate lag length for each variable in each equation. This procedure is followed because there is little basis in economic theory to believe that the same length is appropriate for all variables in each equation. Furthermore, when a common lag length is employed, this common lag length must be kept generous in order to avoid underspecifying the lag for one or more variables and thereby avoiding biased coefficent estimates. Lengthening the common lag by one increases the number of parameters by the square of the number of variables; thus, extensions in lag length rapidly deplete the degrees of freedom for estimation. The alternative approach used here allows a potential reduction in the number of parameters to be estimated.

The use of VARs to analyze the macro effects of deficits is motivated in part by Fischer's (1981) observation that this technique allows one to capture empirical regularities in the data and to thereby gain insight into the channels through which deficits operate. Sims (1982, p. 138) further noted that "careful attention to the historical data exerts an important discipline on what can be plausibly asserted about the way the economy works." However, as is well known, VARs are a reduced-form technique; thus, it is often difficult, based upon the VAR results, to distinguish sharply among structural hypotheses. Some uses of VARS have recently been examined critically by Cooley and LeRoy (1985) and Leamer (1985). Although they are critical of many common uses of VARs, Cooley and LeRoy note, as does Eichenbaum (1985), that there are valid uses of VARs. These include forecasting, the description of the cyclical behavior of a system, the generation of stylized facts about the behavior of the elements of the system which can be compared with existing theories or can be used in formulating new theories, and the testing of theories which generate Granger-causality implications. We regard our investigation as in the spirit of searching for empirical regularities among the macro variables in our system. To achieve this end, we examine the patterns of Granger-causality in the data and compute variance decompositions.

The specification of the VAR precedes equation by equation, and, as mentioned earlier, the *FPE* criterion is used to determine the appropriate lag length for each variable. In specifying the system, logarithms of money, prices, and industrial production are employed, and the actual series for the interest rate and the deficit are used. Since the use of the *FPE* criterion requires stationary data, a trend term was included in each equation. However, as Nelson and Plosser (1982) have pointed out, if a variable follows a difference stationary process, then differencing the variable is the appropriate method of handling mean nonstationarity. Based on this consideration, the sensitivity of the results to inducing stationarity by including a linear trend was checked by specifying the system using the first differences of the logs of money, prices, and industrial production and the first differences of the levels of the interest rate and the deficit. Since the likelihood ratio tests and the variance decompositions for both systems were quite similar, only the results for the system with the linear trend are reported.

The specification procedure is illustrated with reference to the industrial production equation. In order to simplify the presentation, the trend term and a dummy variable designed to capture the effects of the Hawley-Smoot tariff (to be described in the next section) are omitted from the illustration (but are, of course, explicitly considered in specifying the model's equations). The first step is the determination of the lag length for industrial production. This is done by varying the lag in the autoregression $LIP_t = a_0 + a_1(L)LIP_t + e_t$ from 1 to n where LIP_t = log of industrial production, $a_1(L)$ is a distributed lag polynomial such that

$$a_1(L)=\sum_{k=1}^{n} a_{1k}L^k, \tag{1}$$

L is the lag operator so that $L^kLIP_t = LIP_{t-k}$, n = highest order lag,[2] and e_t = zero mean white-noise error term. The FPE is calculated for each autoregression and is defined for lag k, $k = 1, \cdots$, n as

$$FPE_{(k)}=[(T+k+1)/(T-k-1)] \cdot (SSR_{(k)}/T) \tag{2}$$

where T = number of observations used in estimating the autoregression, and SSR = sum of squared residuals. The lag length that minimizes the *FPE* is selected as the order of $a_1(L)$.

Hsiao (1981) pointed out that the *FPE* criterion is equivalent to using an F-test with a varying significance level. As Judge et al. (1982) noted, an intuitive reason for using the *FPE* is that an increase in the lag length increases the first term but decreases the second term and these opposing forces are balanced when their product reaches a minimum. Thus, according to Hsiao (1981, p. 88), the *FPE* criterion is "appealing because it balances the risk due to the bias when a lower order is selected and the risk due to the increase of variance when a higher order is selected."

Once the order of $a_1(L)$ is found, bivariate equations of the following type are estimated for each of the other variables under consideration:

$$LIP_t=a_0+a_1(L)LIP_t+a_2(L)X_t+e_t \tag{3}$$

where $a_2(L)$ is a distributed lag polynomial defined in a similar manner to $a_1(L)$, and X_t = other variables in the system (considered one at a time). $a_1(L)$ is fixed at its previously determined order (k), and the lags in $a_2(L)$ are varied over l, $l = 1, \cdots$, n. The *FPE*s for the resulting equations are defined for lag l, $l = 1, \cdots$, n as

$$FPE_{(k,l)}=[(T+k+l+1)/(T-k-l-1)] \cdot (SSR_{(k,l)}/T). \tag{4}$$

The lag length for X_t that yields that minimum *FPE* is selected as the lag order for that variable.

The next step is the estimation of trivariate equations involving the lagged values of *LIP* and lagged values of two of the other variables under consideration. A problem emerges at this point since the specification of the equation within which the Granger-causality testing will be performed is not, in general, invariant to the order in which the variables are added to the equation. A particular criterion—the specific gravity criterion of Caines, Keng, and Sethi (1981)—is used to determine the order in which the other variables are added to the equation. The specific gravity of *LIP* with respect to, for example, the deficit is defined as the reciprocal of the *FPE* in the bivariate *LIP*-deficit equation. The specific gravity of *LIP* with respect to the other variables is defined analogously. These variables are ranked in order of decreasing specific gravity. The variable with the highest specific gravity is added to the *LIP* equation with the lag order from the relevant bivariate equation.

[2] An n = 12 was predetermined.

Trivariate equations for the remaining variables are estimated, the *FPE*s are calculated, and the variables are ranked in order of their specific gravities. The variable with the highest specific gravity is added to the equation, and the procedure continues until all variables are added to the equation. At this point we have an equation that contains lagged values of *LIP* as well as lagged values of the other variables in the system.

This procedure is repeated for each equation; the resulting equations are combined to form the system. In the VAR estimated here only lagged values of the system's variables appear as right-hand-side variables in the system's equations. Following Hsiao (1981) and Caines, Keng, and Sethi (1981), it is assumed that any contemporaneous relationships are reflected in correlation of error terms across the system's equations. Based on this assumption, the seemingly unrelated regression technique is used to estimate the system.

III. Empirical Results

As noted earlier, the VAR contains five variables—deficits, money, prices, industrial production, and the interest rate. The deficit measure (*DEFF*) is in billions of dollars and is taken from Firestone (1960, Table A-3) and is calculated on a cash basis as federal *receipts minus expenditures.* When DEFF is positive, the actual government budget is in surplus. The money variable is also in billions of dollars and is Friedman and Schwartz' (1963) M2 measure from their Table A-1. The price variable is the wholesale price index (WPI) (base year 1926) and is taken from the 1933, 1938, and 1943 editions of the *Statistical Abstract of the United States.* Industrial production (IP) data (base year 1977) are taken from the 1985 revision of *Industrial Production* (Board of Governors of the Federal Reserve System 1985). The interest rate series (RCP) is the 4–6 month prime commercial paper rate taken from *Banking and Monetary Statistics 1914–1941* (Board of Governors of the Federal Reserve System 1943, Table 120). All data with the exception of RCP are seasonally adjusted.[3]

Based on the procedures described in the previous section, the following model was specified and estimated using seemingly unrelated regression and monthly data over the

[3] The use of seasonally adjusted data was necessitated by the lack of a reliable seasonally unadjusted series for money. Burbidge and Harrison (1985) constructed a seasonally unadjusted series for M1 (currency plus demand deposits), but their series contains deposit data only for member banks of the Federal Reserve in 101 leading cities. The series in Table A-1 (Friedman and Schwartz 1963) are all seasonally adjusted, and since the coverage of these series is more comprehensive than that of Burbidge and Harrison it was decided to use the M2 series seasonally adjusted. A seasonally adjusted deficit series was constructed by seasonally adjusting federal expenditures and receipts from Table A-3 of Firestone (1960) and subtracting expenditures from receipts. The X-11 procedure was used to seasonally adjust expenditures and receipts and was also used to seasonally adjust WPI. Finally, the seasonally adjusted deficit series constructed by Firestone was used in place of the series adjusted by X-11, but the results were not substantially affected.

As a referee pointed out, the use of seasonally adjusted data may, under certain circumstances, spuriously indicate Granger-causality from one variable to another since the temporal ordering of the variables may be obscured by the use of a two-sided moving average procedure such as the X-11 method to seasonally adjust the variables. The referee suggested the following possibility. Suppose that the deficit does not actually Granger-cause output and that output has contemporaneous and lagged effects on the deficit. Assume further that both output and the deficit are serially correlated. Using a two-sided moving average to seasonally adjust each series might, in this case, make it appear that the deficit Granger-causes output. Since we do not find strong evidence of Granger-causality from the deficit to any of the system's other variables, the use of seasonally adjusted data does not seem to generate this type of problem in this study.

sample period July 1922 to June 1938:

$$\begin{bmatrix} LIP \\ LWPI \\ RCP \\ LM2 \\ DEFF \end{bmatrix} = \begin{bmatrix} a_{11}^{9}(L) & a_{12}^{11}(L) & a_{13}^{1}(L) & a_{14}^{7}(L) & a_{15}^{1}(L) \\ a_{21}^{3}(L) & a_{22}^{8}(L) & a_{23}^{9}(L) & a_{24}^{2}(L) & a_{25}^{1}(L) \\ a_{31}^{1}(L) & a_{32}^{1}(L) & a_{33}^{4}(L) & a_{34}^{1}(L) & a_{35}^{1}(L) \\ a_{41}^{2}(L) & a_{42}^{1}(L) & a_{43}^{4}(L) & a_{44}^{7}(L) & a_{45}^{1}(L) \\ a_{51}^{8}(L) & a_{52}^{1}(L). & a_{53}^{1}(L) & a_{54}^{2}(L) & a_{55}^{3}(L) \end{bmatrix} \begin{bmatrix} LIP \\ LWPI \\ RCP \\ LM2 \\ DEFF \end{bmatrix}$$

$$+ \begin{bmatrix} a_1 \\ a_2 \\ a_3 \\ a_4 \\ a_5 \end{bmatrix} + \begin{bmatrix} d_1 \\ d_2 \\ d_3 \\ d_4 \\ d_5 \end{bmatrix} [HSD] + \begin{bmatrix} t_1 \\ t_2 \\ t_3 \\ t_4 \\ t_5 \end{bmatrix} [TIME] + \begin{bmatrix} e_1 \\ e_2 \\ e_3 \\ e_4 \\ e_5 \end{bmatrix}. \tag{5}$$

LIP, *LWPI*, and *LM2* are the logarithms of *IP*, *WPI*, and *M2*, respectively. The a_{ij}, $i = 1, \cdots, 5$; $j = 1, \cdots, 5$, are polynomials in the lag operator L; the superscripts represent the maximum power of L and hence the length of the lag. The a_i, $i = 1, \cdots, 5$, are constant terms, and the d_i, $i = 1, \cdots, 5$, are coefficients on the Hawley-Smoot tariff dummy variable (defined momentarily). The t_i, $i = 1, \cdots, 5$, are coefficients on the time trend variable (*TIME*), and the e_i, $i = 1, \cdots, 5$, are error terms. The individual coefficient estimates are not presented here in order to conserve space, but are available on request. Box-Pierce Q-statistics indicated that for every equation the residuals were free from serial correlation. All estimation was done using the micro version of RATS.

Since many students of the Great Depression attribute a substantial role for the Hawley-Smoot tariff of June 1930 in worsening the recession (see, for example, Meltzer 1976), a crude attempt was made to control for the effects of this tariff. A dummy variable (*HSD*), which consists of zeros from the beginning of the sample to May 1930, ones from June 1930 to May 1934, and a linear decline in value below one from June 1934 to the end of the sample, was added to each equation. The last element of the dummy variable is designed to capture the growing importance over time of the reciprocal trade agreements that ameliorated the effects of the original tariff. Over time more of these reciprocal agreements were reached and this should have reduced the impact of the tariff on the macroeconomy. A likelihood ratio test of the hypothesis that the coefficients on *HSD* jointly equal zero easily rejected this hypothesis ($\chi^2 = 30.22$ at a marginal significance level of 1%).[4] The coefficients on *HSD* are negative in the *LIP*, *LWPI*, *RCP*, and *DEFF* equations. The *HSD* coefficient in the *LWPI* equation is highly significant (1% level) and is significant at the 5% level in the *RCP* equation. The

[4] The likelihood ratio statistic is computed as $-2 \log (L^c/L^u)$ where L^c is the maximized likelihood of the constrained system (the system in which the HSD terms have been dropped) and L^u is the maximized likelihood of the unconstrained system. This statistic asymptotically follows a chi-square distribution with n degrees of freedom, where n is the number of imposed constraints.

A likelihood ratio test also revealed that the hypothesis that the coefficients on time jointly equalled zero could also be rejected.

Table 1. Likelihood Ratio Tests

	Hypothesis	Likelihood ratio statistic[a]
a.	$a^1_{15}(L) = a^1_{25}(L) = a^1_{35}(L) = a^1_{45}(L) = 0$	6.50 (.16)
b.	$a^1_{15}(L) = 0$	0.84 (.36)
c.	$a^1_{25}(L) = 0$	4.55 (.03)
d.	$a^1_{35}(L) = 0$	0.91 (.34)
e.	$a^1_{45}(L) = 0$	0.10 (.75)
f.	$a^{11}_{12}(L) = a^1_{32}(L) = a^1_{42}(L) = a^1_{52}(L) = 0$	67.19 (.00)
g.	$a^{11}_{12}(L) = a^1_{15}(L) = a^1_{32}(L) = a^1_{35}(L) = a^1_{42}(L) = a^1_{45}(L) = 0$	67.84 (.00)
h.	$a^{11}_{12}(L) = 0$	44.17 (.00)
i.	$a^1_{32}(L) = 0$	5.81 (.02)
j.	$a^1_{42}(L) = 0$	22.52 (.00)
k.	$a^1_{52}(L) = 0$	0.37 (.55)
l.	$a^8_{51}(L) = a^1_{52}(L) = a^1_{53}(L) = a^2_{54}(L) = 0$	46.83 (.00)
m.	$a^8_{51}(L) = 0$	30.07 (.00)
n.	$a^1_{53}(L) = 0$	0.19 (.67)
o.	$a^2_{54}(L) = 0$	5.50 (.06)

[a] The marginal significance level is in parentheses beside the computed statistic.

coefficient is significant at the 10% level in the *LIP* equation, but is not significant in the *DEFF* equation. The coefficient on *HSD* in the *LM2* equation is positive but not significant at standard levels. On balance, the dummy variables provide evidence of a significant effect of the Hawley-Smoot tariff with a theoretically expected direction of effect on *LIP, LWPI,* and *RCP.*

The role of the deficit in the system was checked by a sequence of likelihood ratio tests. As Granger (1969) proved, a zero-element in the off-diagonal elements of a system like (5) indicates the absence of direct Granger-causality from one variable to another. The results of the likelihood ratio tests for DEFF are presented in Table 1. Hypothesis a is a *joint* test of the hypothesis that *DEFF* has no significant effects on the other variables in the system. The computed likelihood ratio statistic is not significant at even the 10% level so that hypothesis a cannot be rejected. Hypotheses b through e are bivariate Granger-causality tests for *DEFF.* These tests indicate no direct Granger-causality from *DEFF* to *LIP* (hypothesis b), *RCP* (hypothesis d), and *LM2* (hypothesis e). However, the rejection of hypothesis c indicates that *DEFF* directly Granger-causes *LWPI.* Although the joint test indicates no effects of *DEFF* within the system, the bivariate test of hypothesis c indicates Granger-causality from *DEFF* to *LWPI.* This can arise because of the difference between joint and individual hypothesis tests. For a discussion of this, see Kennedy (1985). Similar results were found for the system in first-differences.

Since *DEFF* directly Granger-causes *LWPI,* DEFF will indirectly Granger-cause the other variables in the system if *LWPI* Granger-causes these variables. Hypothesis f is a joint test of the hypothesis that *LWPI* has no significant effects on the other variables in the system. The computed likelihood ratio statistic indicates that this hypothesis can be rejected. Hypothesis g is a joint test of the hypothesis that *DEFF* and *LWPI* have no effects on the other system variables. Because of the bivariate Granger-causality from *DEFF* to *LWPI,* DEFF and *LWPI* are treated as a block in this hypothesis. The hypothesis can be easily rejected. Hypotheses h through k are bivariate Granger-causality tests for *LWPI*. These test indicate bivariate Granger-causality from *LWPI* to *LIP* (hypothesis h), *RCP* (hypothesis i), and *LM2* (hypothesis j), but no Granger-causality from *LWPI* to *DEFF* (hypothesis k). Taken together, hypothesis tests f through j indicate indirect Granger-causality from *DEFF* to *LIP, RCP,* and *LM2*. Again, the same pattern of significant and insignificant likelihood ratio statistics was found when the same tests were performed for the first-difference system.

The effects of the other system variables on *DEFF* are also presented in Table 1. Hypothesis *l* is a joint test of the hypothesis that the other variables in the system have no significant effect on *DEFF*. This hypothesis is easily rejected. Hypotheses m through o and k are bivariate Granger-causality tests. We see that *LIP* (hypothesis m) and *LM2* (hypothesis o) Granger-cause *DEFF* but that *LWPI* and *RCP* do not.

The likelihood ratio tests indicate that *DEFF may* have important economic effects, because *DEFF* appears to directly Granger-cause *LWPI* and, through its effects on *LWPI*, to indirectly Granger-cause the other variables. However, as Sims (1972) has noted, the absolute sizes of the coefficients on a variable like *DEFF* are important regardless of what the likelihood ratio tests indicate. He points out that coefficients that are statistically significant may be so small that they are not economically significant. Sims (1982) has more recently suggested that VDCs can be used to measure the strength of Granger-causal relations. VDCs show the proportion of the forecast error variance for each variable that is attributable to its own innovations and to shocks to the other system variables. Both *direct* and *indirect* effects are captured in the VDCs. The VDCs are based upon the moving average representation of the system and the moving average representation reflects the sizes of the estimated coefficients. As Sims noted (p. 131), "A variable that is optimally forecast from its own lagged values will have all its forecast error variance accounted for by its own disturbances." Thus if *DEFF* explains only a small portion of the forecast error variance of *LWPI*, this could be interpreted as a weak Granger-causal relation.

Variance decompositions for system (5) are generated in the manner suggested by Sims (1980). This method recognizes that, in general, the correlation of residuals across equations is not zero. In calculating the VDCs the variables are ordered in a particular fashion. Because of the cross-equation residual correlation, when a variable higher in the order changes, variables lower in the order are assumed to change. The extent of the change depends upon the covariance of the variables higher in the order with those lower in the order. Because of this the VDCs may be sensitive to the ordering of the variables so that it is useful to examine several orderings.

Bernanke (1986) noted that the use of VDCs based upon a variance-covariance matrix orthogonalized by the Choleski decomposition as we have done imposes a recursive structure upon the model and that this is defensible if one believes that the model is, indeed, recursive. Examination of different orderings can be justified if one believes the model to be recursive but is uncertain about the causal ordering of the variables. In terms of Bernanke's argument, we assume a block-recursive model with the financial sector block (*DEFF, RCP, LM2*) preceding the goods market block (*LIP, LWPI*). We contend that financial market data are more quickly available and that financial markets clear more rapidly than the goods market so that is is reasonable to order the financial variables before the goods market variables; nevertheless, among the financial variables it may be more difficult to justify any particular ordering. Thus, we experiment with several orderings of the financial variables.

We initially choose an ordering (*DEFF, RCP, LM2, LIP, LWPI*) based on the following theoretical considerations. A number of theories of how deficits can affect the money supply have been advanced in the professional literature and in the popular press. Perhaps the most common story assigns a crucial institutional role to the Federal Reserve. Budget deficits are said to drive up interest rates and crowd out private expenditures unless the Federal Reserve monetizes at least part of the newly issued government debt. The Fed may thus be enticed (or coerced) into open market purchases designed to mitigate the upward pressure on interest rates, and this will lead to more rapid growth in the money supply. Faster money growth can, in the short run at least, expand real output, and will surely lead to higher inflation, although with perhaps a longer lag. But, as noted earlier, we also experiment with other orderings of the financial variables as follows: *DEFF, LM2, RCP*; *LM2, DEFF, RCP*; and *LM2*, RCP, DEFF.

Variance decompositions for the ordering *DEFF, RCP, LM2, LIP, LWPI* are shown for different months over a 48-month horizon in column 1, Table 2. Results for the deficit are virtually identical when alternative orderings of the financial market variables are chosen.[5] Our results indicate that deficits have a weak effect. There is little or no support for the view that deficits are either monetized or have significant interest rate effects, nor is there indication of substantial effects on the goods market. While deficits Granger-cause prices, the percentage of forecast error variance in *LWPI* attributable to *DEFF* is at most around 7%. The effects of *DEFF* in the first-difference system are somewhat weaker than those reported in Table 2, column 1.

The deficit is mainly accounted for by its own innovations. While both production and money Granger-cause the deficit, both relations are quite weak.

IV. An Alternative Specification

In this section we examine an alternative specification of the VAR that uses a common lag length for each variable in each equation, as was done by Sims (1980) and Burbidge and Harrison (1985). Our motivation is primarily twofold. First, we want to check the sensitivity of our results to the technique used to estimate the VAR. Second, there is the question of structural instability. Because of the effort involved in specifying Hsiao-type VARs, the structural instability issue is addressed using the Sims-type VAR, which can be specified more quickly and easily. Burbidge and Harrison reported that for their interwar period a likelihood ratio test revealed significant problems when they estimated their VAR for two subperiods, the first ending in October 1929 and the second beginning the following month.[6] We thus estimate separate VARs for the July 1922–October 1929 and November 1929–June 1938 subperiods.

Following Lutkepohl (1982), Akaike's AIC criterion is used to determine the lag length of the Sims-type VARs. The lag length chosen is the one that minimizes

$$AIC(k) = \ln \det\Sigma_k + 2d^2/T, \; k = 1, \cdots, m$$

where d = the number of variables in the system, m = maximum lag length considered,

[5] Complete results for the entire system for both the system with the linear trend and the system in first differences for various orderings of variables, as well as for Granger-causality tests, are available on request.

[6] Burbidge and Harrison (1985, p. 48, fn. 5).

Table 2. Variance Decompositions Explained by Innovations in DEFF

Relative variation in	Months later	22,7-38,6 variable lag	22,7-38,6 common lag = 2	22,7-29,10 common lag = 12	29,11-38,6 common lag = 2
DEFF	1	100.0	100.0	100.0	100.0
	6	93.6	96.2	45.2	94.6
	12	89.5	95.5	31.4	94.2
	24	86.0	94.6	26.1	93.4
	36	85.1	93.9	24.1	92.9
	48	84.9	93.8	23.6	92.8
RCP	1	0.1	0.2	0.0	0.4
	6	0.3	0.0	3.0	0.4
	12	0.5	0.1	4.8	0.5
	24	0.6	0.3	4.4	0.9
	36	0.9	0.9	4.8	2.5
	48	1.2	1.3	5.8	2.9
LM2	1	1.0	1.0	0.0	1.6
	6	3.0	5.1	5.4	11.1
	12	5.3	7.7	10.7	16.0
	24	7.2	9.4	12.7	17.0
	36	7.6	9.4	11.7	16.2
	48	7.5	9.2	11.2	16.2
LIP	1	0.3	1.3	5.0	1.0
	6	2.3	10.6	1.6	11.6
	12	3.2	13.5	1.4	14.6
	24	4.1	13.4	2.1	14.3
	36	3.9	13.3	1.9	14.3
	48	3.9	13.2	2.7	14.4
LWPI	1	1.1	1.0	1.0	1.4
	6	5.8	9.6	3.3	13.0
	12	6.7	12.0	6.6	15.5
	24	7.1	11.4	7.6	14.8
	36	6.8	11.1	5.8	14.8
	48	6.9	11.2	5.8	15.1

$det\Sigma_k$ = determinant of Σ_k, Σ_k = diagonal matrix of the estimated residual variance for each equation in the system. Use of the AIC criterion suggested a common lag of 2 months for the full period and for the November 1929–June 1938 subperiod and a common lag of 12 months for the July 1922–October 1929 subperiod. The systems estimated over the full period and the November 1929–June 1938 subperiod included both the trend and the Hawley-Smoot dummy. However, the system estimated over the July 1922–October 1929 subperiod included only the trend term since the Hawley-Smoot dummy variable consisted of zeros over this subperiod.

Variance decompositions using common lag lengths for the interwar period and the two subperiods are also shown in Table 2. While the estimation technique makes some difference for the interwar period as a whole, the deficit variable still has relatively weak effects. The percentage of the forecast error variances of *LIP* and *LWPI* explained by innovations in *DEFF* increases—but only to a maximum of about 14% and 12%, respectively, at the 12-month horizon. *DEFF* still explains virtually none of the forecast error variance in *RCP* and less than 10% in *LM2*. The deficit variable remains free of significant feedback from the other variables.

It might be argued that the macroeconomy should not be considered stable over the entire interwar period and that policy makers did not adhere to a consistent policy rule. Perhaps a regime change occurred at the start of the Great Depression. With respect to the government budget, modest surpluses were the rule prior to November 1929 (*DEFF* is positive) and somewhat larger deficits dominated thereafter. Surpluses occurred in 86 of 88 months in the first subperiod and averaged \$0.76 billion with a variance of 0.16, while deficits occurred in 95 of 104 months in the second subperiod and averaged \$ − 1.92 billion with a variance of 2.76.

Some differences can be noted when the two subperiods are estimated separately and the results compared to each other and to the interwar period as a whole. However, *DEFF* still cannot be considered a dominant variable.

In the first subperiod the percentage of forecast error variance explained by innovations in *DEFF* exceeds 10% (about 13% at the 24-month horizon) only in the case of *LM2*. In the second subperiod, innovations in *DEFF* have slightly stronger effects. Innovations in *DEFF* account for about 17% of the forecast error variance in *LM2* at the 24-month horizon and about 15–16% of the forecast error variances in *LIP* and *LWPI* at the 12-month horizon.

By far the major difference between the two subperiods is the degree to which *DEFF* is explained by its own innovations. In the earlier subperiod, innovations in the other four variables in the system account for a combined 76% of the forecast error variance in *DEFF*. In the later subperiod *DEFF* is explained almost entirely by its own innovations. Conclusions about the deficit variable are not sensitive to the ordering of financial variables.

V. Conclusions

In estimating five-variable VARs for the July 1922–June 1938 period, we first use a technique in which lag lengths differ for each variable in each equation and for the same variable across equations. To check the sensitivity of our result to this technique, we next estimate VARs using a common lag length for each variable in each equation. Since a regime change may have occurred at the start of the Great Depression, we then make separate estimates for two subperiods, July 1922–October 1929 and November 1929–June 1938.

In no case does the ordering of financial variables affect conclusions about the deficit variable. While the impact of the deficit varies slightly across time periods and according to the estimation technique, our deficit variable cannot be considered a dominant variable. There is no evidence of substantial debt monetization or of important effects on interest rates, production, or prices. These results are broadly consistent with those of Joines (for money) and Evans (for interest rates), both of whom found statistically insignificant effects of deficits in somewhat different time periods than ours, and with those studies using post-World War II data that found small or insignificant effects of deficits.

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